Relations industrielles
Industrial Relations

The Union Wage Effect Re-Examined
Le syndicalisme et les salaires : un réexamen de la théorie

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Volume 35, numéro 2, 1980

URI : https://id.erudit.org/iderudit/029060ar
DOI : https://doi.org/10.7202/029060ar

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Éditeur(s)

Département des relations industrielles de l'Université Laval

ISSN

0034-379X (imprimé)
1703-8138 (numérique)

Découvrir la revue

Citer cet article

Maki, D. & Christensen, S. (1980). The Union Wage Effect Re-Examined. *Relations industrielles / Industrial Relations*, *35*(2), 210–230. https://doi.org/10.7202/029060ar

Résumé de l'article

La raison première de cet article provient d'une étude d'Ashenfelter et de Johnson (1972) dans laquelle ils énonçaient ce qui suit: 1° le modèle simple d'équation salariale, typiquement utilisée pour mesurer l'effet des syndicats sur les salaires, est inexact à cause de la présence de variables endogènes parmi les variables explicatives; 2° l'effet présumé des syndicats sur les salaires est de peu d'importance et sans signification quand il est obtenu à partir d'un modèle d'équation simultanée plus précis. Les résultats de notre étude confirment la première proposition, mais ils démontrent la fausseté de la seconde.

Nous avons trouvé une preuve solide de dépendance réciproque entre le taux des salaires et deux de ses variables explicatives, soit la présence de syndicats et la qualité du travail. Cependant, le fait de tenir compte de cette double dépendance ne réduit pas l'effet présumé de l'existence des syndicats à néant. Au contraire, nous persistons à déceler que les salaires conventionnels ont un effet considérable etsignificatif. Les résultats de notre évaluation indiquent que le salaire moyen des travailleurs de la production dans les industries manufacturières est supérieur à 4.2 pour cent en regard d'une différence de 0.10 dans le degré de pénétration syndicale (celle-ci étant la proportion de tous les travailleurs de la production liés par une convention collective de travail). Les résultats que nous avons obtenus indiquent aussi une différence de 51 pour cent entre les salaires conventionnels et les autres, mais l'exactitude de cette estimation dépend, et cela d'une façon décisive, de l'exactitude avec laquelle nous pouvons établir le salaire moyen dans une industrie où il n'y a pas de syndicat. Parce que notre échantillon n'inclut pas de telles industries et ne contient en réalité que trois industries (sur 40) où le taux de syndicalisation est inférieur à 50 pour cent, nous avons peu de confiance dans notre estimation pour la différence d'ensemble.

Les résultats d'une évaluation supplémentaire ont été obtenus dans un effort pour découvrir les raisons d'estimations disparates de l'effet du salaire conventionnel dans la présente étude et dans des études antérieures aux États-Unis et au Canada. Quoique ces résultats ne soient pas bien précis, il semble que les différences dans la spécification du modèle, les industries sélectionnées et l'ampleur de l'échantillon peuvent tous en être responsables. Cependant, les résultats auxquels nous sommes arrivés étaient solides dans toutes les expérimentations variées que nous avons tentées. Nous soupçonnons, bien que nous n'ayons pu le vérifier à partir des données dont nous disposions, que la cause principale de la disparité dans les estimations de l'effet du salaire conventionnel obtenues jusqu'ici réside dans les différents niveaux d'agrégation sur lesquels les diverses évaluations ont été fondées.

The Union Wage Effect Re-Examined

Dennis Maki
and
Sandra Christensen

In this paper, the authors extend the literature on union wage effect in two ways. They use aggregate industry data in a simultaneous-equations model to provide an estimate for the union wage effect in Canada which recognizes the endogenity of union coverages. And they suggest a possible cause for the difference in results obtained for the U.S. by Ashenfelter and Johnson, on the one hand, and Kahn on the other.

In 1963, H.G. Lewis published a comprehensive review of U.S. studies which attempted to measure the extent to which the wage rates of union workers exceeded those of nonunion workers. He concluded that, on average, the wage rates of union workers were 10 to 15 percent above those of nonunion workers, at least for the 1957-58 period. Since that time, a number of additional U.S. studies have been made which have benefited from the more extensive data on labor force and industry characteristics now available. These later studies have yielded a wide range of estimates for the union-nonunion wage differential, but most of them are in excess of the 10 to 15 percent effect suggested by Lewis.[1]

However, in 1972, Orley Ashenfelter and George Johnson published an innovative study of the union wage effect in U.S. manufacturing industries in which they challenge the validity of the single-equation estimation methods used in previous studies. They argue that if the union wage effect is as large as recent studies indicate, it is difficult to understand why so much of the work force remains unorganized. They suggest that the large estimates obtained might be the result of simultaneous equation bias, due to the inaccuracy of the usual model in which union coverage and labor quality are treated as exogenous determinants of wages. They develop an alternative model in which the extent of union coverage, labor quality, and wages are jointly dependent. Their OLS estimate of their wage equation in-

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1 See EVANS and MACDONALD (1976), tables 1 and 2, for a summary of these results.

dicates a large and significant union wage effect. However, when they obtain 2-3SLS estimates, in which the union coverage and labor quality variables in the wage equation are treated as endogenous variables, the estimated union wage effect drops in value and is statistically nonsignificant. This suggests that the usual single-equation result of a large and significant positive union effect on wages gives unions too much credit for raising wages because the variable for union coverage also picks up reverse causality running from wages to union coverage.

In subsequent studies that have adopted the simultaneous-equations approach suggested by Ashenfelter and Johnson, the results for the union effect on wages have been mixed. A study by Kahn (1979), using aggregate U.S. industry data (as did Ashenfelter and Johnson), finds a large and significant union wage effect even when the endogeneity of union coverage is recognized. A study by Schmidt and Strauss (1979), using individual data, finds a small union wage effect which drops to nonsignificance when endogeneity of union coverage is recognized; but studies by Lee (1978) and Leigh (1978), which also use individual data, find the union wage effect remains large and significant.

We are aware of only three Canadian studies which attempt to estimate the union wage effect.[2] Each of these finds a positive and significant union wage effect of about 20 percent, but each uses only the conventional single wage equation approach. In this paper, we extend the literature in two ways: 1) we use aggregate industry data in a simultaneous-equations model to provide an estimate for the union wage effect in Canada which recognizes the endogeneity of union coverage; and 2) we suggest a possible cause for the difference in results obtained for the U.S. by Ashenfelter and Johnson, on the one hand, and Kahn on the other.

A note of caution is in order prior to the presentation of our model and empirical results. The statistical problems inherent in estimating the effect of unions on wages are many. The reader may refer to Reder (1965) for a full discussion of these problems. We outline here just two of the problems which can arise because they may contribute to the disparity between our results and those obtained in some other recent studies.

First, all those studies which, like ours, use aggregate industry data are constrained by the nature of the available wage or earnings statistics. Such data are reported in terms of industry averages, so that the effect of unions on those establishments within the industry which are unionized may be

2 See KUMAR (1972), STARR (1973), and EVANS and MACDONALD (1976).

swamped in the average figure by the presence of many unorganized establishments. This means that there may be considerable variation in the size of the union wage effect estimated in different studies arising solely from differences in the level of industry aggregation in the data used. In this context, our use of three and four digit industries should be preferable to the two-digit industrial classification used in some other studies.

Second, unions may affect wages not only in those establishments which are unionized but also in nonunion establishments which hire in the same labor market. The effect of unions on nonunion wages may be either to raise them, as nonunion employers attempt to forestall unionization (the "threat effect"); or to lower them, as workers displaced by the higher wages in the union sector flood the nonunion sector (the "displacement" effect). Our estimate of the union wage effect is an estimate of the differential which would exist after threat and displacements adjustments in the nonunion sector have taken place. A discussion of the difficulties involved in separating out threat and displacement effects, and an empirical attempt to do so, are contained in the study by Gerald Starr (1973) for the Ontario Ministry of Labor.

SPECIFICATION OF THE MODEL

We estimate the following three-equation model, in which wages, union coverage, and labor quality are endogenous:

(1) $\ln W = a_0 + a_1 U + a_2 ELT9 + a_3 F + a_4 LAGPROF + a_5 ONT + a_6 QUE + a_7 PRA + a_8 BC + e_1$

(2) $U = b_0 + b_1 \ln W + b_2 ELT9 + b_3 BIGEMP + b_4 ONT + b_5 QUE$ $QUE + b_6 PRA + b_7 BC + e_2$

(3) $ELT9 = c_0 + c_1 \ln W + c_2 URBAN + c_3 OLD55 + c_4 ONT + c_5 QUE + c_6 PRA + c_7 BC + e_3.$

Our data set is for 1974, and includes only production workers in Canadian manufacturing industries. The 40 industries included in our sample were all those, for which the necessary data were available.[3]

[3] A complete list of the industries included together with variable definitions, sources, and a listing of the data are included in the attached appendix.

The variables are defined as follows:

W =	average wage for production workers in a given industry, measured empirically as gross earnings of production workers divided by manhours worked;
U =	proportion of production workers covered by a collective agreement;
ELT9 =	proportion of production workers with education less than grade 9 (1971);
F =	proportion of production workers that are female;
LAGPROF =	percent profit before tax (1973);
BIGEMP =	proportion of workers in establishments employing 200 or more;
URBAN =	proportion of workers residing in Census Metropolitan Areas (1971);
OLD55 =	proportion of workers who are 55 or older (1971);
ONT =	proportion of establishments located in Ontario;
QUE =	proportion of establishments located in Quebec;
PRA =	proportion of establishments located in the Prairie provinces;
BC =	proportion of establishments located in British Columbia.

THE WAGE EQUATION (1)

The specification for the equation explaining average industry wage can be justified by viewing it as the reduced-form equilibrium relationship resulting from a simple labor market model. Hence, its arguments should be the nonwage determinants of labor demand and supply. Consequently, we include variables for union coverage (U), since the existence of a collective bargaining relationship can be expected to affect the offer price of labor; for the proportion of workers who are female (F), since this could pick up either demand effects based on employer discrimination or supply effects based on male-female differences in offer price; for the industry's profit rate over the previous year (LAGPROF), since this is indicative of the employer's "ability to pay" and may also indicate increased demand for labor due to industry expansion; and for labor quality (ELT9), since this affects productivity and hence the demand for labor. In the empirical specification for labor quality, it is customary to use a measure for average education. We use the proportion of workers who have not completed grade 9. We think this is conceptually superior to average education because, in our sample of production workers in manufacturing, the main effect of education on earnings is probably as an indicator of trainability. In Canada, completion of grade 10 is a prerequisite for entry into many apprenticeship programs, and hence not completing grade 10 would have been the most appropriate measure. The grade 9 measure is the closest we could come with available data.

Equation (1) is similar to those estimated in previous studies of the union wage effect, using OLS. However, the equation includes some variables for which arguments of reverse causation can be made, making OLS estimation biased, even asymptotically. In particular, greater union coverage may not only raise the average wage in an industry; it may also be that higher wage industries are more likely to be unionized, since workers in such industries have more to protect. Ashenfelter and Johnson (1972, pp. 491-492) explain this as the result of increased demand by high wage (and hence high income) workers for all normal goods, including union services. Further, although a relatively high quality labor force in an industry necessitates relatively high wages, it may also be that a high wage industry will attract a relatively high quality labor force since it can choose the "cream of the crop" among job seekers. Because of the existence of this two-way causation between dependent and independent variables in the wage equation, we must go on to specify equations for union coverage and labor quality in order to obtain consistent estimates from the model.

THE UNION COVERAGE EQUATION (2)

As before, an appropriate specification for this equation can be obtained by viewing it as the reduced-form equilibrium relationship resulting from a simple demand and supply model for union services. As indicated earlier, we would expect the workers' demand for union services to vary directly with the industry wage level (ln W), since workers in high wage industries have more to protect; and with our inverse measure of education (ELT9), since the least-educated or least-skilled workers have the least potential for gains from individual bargaining. We would expect the supply of union services to vary directly with the size of potential bargaining units within the industry (BIGEMP), since this reduces the cost-to-benefit ratio of organizational efforts.

THE LABOR QUALITY (Education) Equation (3)

As indicated earlier, we expect our inverse education measure to be negatively related to the average industry wage (ln W), since employers with relatively attractive pay scales can be very selective in hiring, thus building up a relatively high quality labor force. Further, we would expect other determinants of the educational mix of an industry's labor force to include the age of the labor force (OLD55), since near-universal education at least through high school is a post-World War II phenomenon in Canada; and urban location (URBAN), since the average educational attainment of the population tends to be higher in urban than rural areas. Finally, we include regional controls in this and the other two equations. The Atlantic provinces serve as the reference group, with controls explicitly entered for Ontario (ONT), Quebec (QUE), the Prairie provinces (PRA), and British Columbia (BC).

EMPIRICAL RESULTS

Our estimation results for the three equations are presented in Tables 1, 2, and 3. Our primary interest is in the coefficient attached to the union coverage variable in the wage equation. From Table 1, it is clear that we do *not* find that the estimated union wage effect drops to nonsignificance once the endogeneity of union coverage in the wage equation is recognized. In fact, there is no significant difference in the estimates we obtain for the coefficient of U, whether the estimates are obtained from OLS; from 2-3SLS in a two-equation model where wages and union coverage are jointly determined; or from 2-3SLS in a three-equation model where labor quality is added as a third jointly determined variable. While there is evidence, in Table 2, of significant reverse causality running from wages to union coverage, allowance for this reverse causality in our estimation procedure does not destroy or even significantly weaken the causal relationship running from union coverage to wages.

Using the approach first developed by Lewis (1963), an estimate of the union-nonunion wage differential, defined as:

$$d = (W_u - W_n)/W_n = (W_u/W_n) - 1,$$

can be derived from the coefficient attached to the union coverage variable in the wage equation.[4] Since equation (1) says that:

$$\ln W = a_0 + a_1 U + a X$$

4 This assumes the differential is constant across industries.

where X = all other explanatory variables for ln W, it follows that:

$$W = e^{a_1 U} e^{K}$$

where $K = a_0 + a X$. Hence:

$$W_u = e^{a_1} e^{K} \text{ since } U = 1 \text{ for a completely unionized industry;}$$

$$W_n = e^{O} e^{K} \text{ since } U = 1 \text{ for a nonunionized industry;}$$

and $d = (W_u / W_n) - 1 = e^{a_1} - 1.$

Our three-equation 3SLS estimate of a_1 is .412, which implies an estimated union-nonunion wage differential of 51 percent. However, it should be noted that our calculated value for W_n, the average wage in an industry with no union coverage, involves extrapolation of our results far outside the range of union coverage values contained in our sample.[5] Hence, we can have little confidence in the value calculated for W_n, and therefore our confidence in the accuracy of the large union-nonunion wage differential we have obtained is also very low. Similar problems exist with virtually all aggregate industry studies of this type. This is probably a large part of the explanation for the wide range of estimates reported for the union-nonunion differential, even from single-equation studies.

However, our results can be used with more confidence to assess the difference in average industry wage to be expected when comparing one highly unionized industry to another which differs only in being slightly more (or less) unionized. For example our results indicate that, for two otherwise identical industries, a 0.10 difference in the proportion of workers covered by a union contract would be accompanied by an average wage differential of 4.2 percent.

Our results contradict those obtained for the U.S. by Ashenfelter and Johnson, since our results show that the effect of unions in raising average industry wages remains large and significant even after the endogeneity of union coverage and labor quality is accounted for in the estimation procedure. However, our results are very similar to those reported by Kahn, who obtained a highly significant 3SLS estimate of..55 for the coefficient of

5 The lowest value for U in our sample is .17; only 3 of the 40 industries in our sample have U values less than .50. At the other extreme, 10 out of the 40 have U values between .90 and .99.

TABLE 1

Estimated Regression Coefficients for Wage Equation

Estimation Procedure	*Dependent Variable*	*Coefficients of* *U*	*ELT9*	*F*	*LAGPROF*	*ONT*	*QUE*	*PRA*	*BC*	*CONST*	R^2	*SEE*
OLS	ln W	.455**	-.646**	-.499**	.009*	-.026	.348*	.389	.512	1.190**	.896	.084
		(.095)	(.197)	(.087)	(.004)	(.163)	(.158)	(.217)	(.280)	(.205)		
2 eqn. 2SLS	ln W	.533**	-.664**	-.482**	.009*	-.040	.381*	.404	.511	1.128**	.894	.085
		(.119)	(.200)	(.089)	(.004)	(.165)	(.162)	(.220)	(.283)	(.215)		
2 eqn 3SLS	ln W	.533**	-.674**	-.491**	.007*	-.049	.384*	.398	.504	1.147**		
		(.119)	(.199)	(.088)	(.004)	(.164)	(.162)	(.220)	(.283)	(.213)		
3 eqn 2SLS	ln W	.452**	-1.696**	-.377**	.003	-.539	.205	-.147	.156	1.919**	.800	.117
		(.166)	(.529)	(.131)	(.007)	(.315)	(.236)	(.387)	(.419)	(.456)		
3 eqn 3SLS	ln W	.412**	-1.779**	-.348**	.006*	-.568	.166	-.199	.134	1.985**		
		(.111)	(.502)	(.093)	(.003)	(.304)	(.232)	(.379)	(.415)	(.432)		

** Significant at the .01 level
* Significant at the .05 level

Standard errors are in parentheses. The 2 equation model does not include the equation for labor quality (ELT 9). Tests for the coefficients of U, ELT9, F, and LAGPROF are 1-tailed; others are 2-tailed since no sign was hypothesized. The set of regional controls is significant at the .05 level.

TABLE 2

Estimated Regression Coefficients for Union Coverage Equation

Estimated Procedure	*Dependent Variable*	Coefficients of *lnW*	*ELT9*	*BIGEMP*	*ONT*	*QUE*	*PRA*	*BC*	*CONST*	R^2	*SEE*
OLS	U	.356**	.608**	.428**	.127	-.387*	-.100	-.248	-.124	.802	.088
		(.109)	(.217)	(.071)	(.167)	(.154)	(.234)	(.303)	(.256)		
2 eqn 2SLS	U	.353**	.605**	.429**	.127	-.387*	-.098	-.245	-.119	.802	.088
		(.138)	(.231)	(.075)	(.167)	(.154)	(.238)	(.314)	(.286)		
2 eqn 3SLS	U	.353**	.605**	.429**	.127	-.387*	-.098	-.245	-.119		
		(.138)	(.231)	(.075)	(.167)	(.154)	(.238)	(.314)	(.286)		
3 eqn 2SLS	U	.472*	1.140*	.414**	.350	-.338	.099	-.150	-.607	.764	.096
		(.221)	(.656)	(.085)	(.300)	(.177)	(.315)	(.341)	(.658)		
3 eqn 3SLS	U	.499*	1.256*	.418**	.396	-.326	.141	-.135	-.718		
		(.214)	(.640)	(.077)	(.296)	(.177)	(.313)	(.341)	(.642)		

** Significant at the .01 level
 * Significant at the .05 level

Standard errors are in parentheses. The 2 equation model does not include the equation for labor quality (ELT9). Tests for ln W, ELT9 and BIGEMP are 1-tailed; others are 2-tailed since no sign was hypothesized. The set of regional controls is significant at the .01 level. Because the wage equation is just identified in our two-equation specification, there is no gain in efficiency of estimation for the union coverage equation as we move from 2SLS to 3SLS. See A.S. GOLDBERGER (1964), p. 352.

TABLE 3

Estimated Regression Coefficients for Labor Quality Equation

Estimation Procedure	*Dependent Variable*	*Coefficients of* *ln W*	*URBAN*	*OLD55*	*ONT*	*QUE*	*PRA*	*BC*	*CONST*	R^2	*SEE*
OLS	ELT9	-.214**	-.126*	.593	-.239	.022	-.224	-.067	.788*	.703	.067
		(.067)	(.072)	(.464)	(.133)	(.129)	(.183)	(.236)	(.152)		
3 eqn 2SLS	ELT9	-.247**	-.127*	.569	-.224	.019	-.189	-.014	.825**	.701	.067
		(.074)	(.073)	(.466)	(.134)	(.129)	(.186)	(.242)	(.156)		
3 eqn 3SLS	ELT9	-.250**	-.129*	.395	-.235	.013	-.188	-.009	.860**		
		(.073)	(.068)	(.240)	(.134)	(.129)	(.183)	(.240)	(.135)		

** Significant at the .01 level
* Significant at the .05 level

Standard errors are in parentheses. Tests for ln W, URBAN, and OLD55 are 1-tailed; others are 2-tailed since no sign was hypothesized.

the union coverage variable. Kahn (1979, p. 528) speculates that the difference between his results and those of Ashenfelter and Johnson is caused by Kahn's inclusion of a fourth endogenous variable, production technique (SK), in the model. However, this seems unlikely to be the explanation, for two reasons: *first*, because our results are similar to Kahn's although we use only a three-equation model similar to that of Ashenfelter and Johnson; and *second*, because Kahn's estimation results show that a) he fails to find any variables which contribute significantly to the explanation of his production technique variable (SK); b) SK is not a significant variable in any of the other three equations in his model. Hence, the reasonable conclusion would seem to be that the addition of that fourth endogenous variable to the three-equation model used by Ashenfelter and Johnson contributes nothing to the results.

We tentatively suggest an alternative explanation for the difference between both Kahn's results and ours, on the one hand, and those of Ashenfelter and Johnson on the other. It appears that the model used by Ashenfelter and Johnson is underspecified, given the significant coefficients estimated for the additional explanatory variables we have included in the equations. Further, Ashenfelter and Johnson have only 19 (two-digit) industries in their sample. By contrast, our sample includes 40 (three-digit) industries, and Kahn's sample includes 49 (three-digit) industries. It is well known that simultaneous-equation estimation techniques are very sensitive to specification error. That, together with the small degrees of freedom in estimation, make it unsurprising that the results obtained by Ashenfelter and Johnson are unstable and weak.

ADDITIONAL RESULTS

In an attempt to investigate further some of the sources of instability between studies in the estimation of the union-nonunion wage differential, three supplementary estimations were performed. In the first of these, we attempted a rough test of the explanation we suggested at the end of the preceding section for the weak and unstable results obtained by Ashenfelter and Johnson. The test was made in three parts. First, we arbitrarily divided our sample of 40 industries into two samples of 20 industries each by selecting every other industry from the SIC list, and we estimated our wage equation (1) for each of these two sub-samples. Second, we reduced the specification of our three-equation model to one which was as nearly as possible comparable to that used by Ashenfelter and Johnson. In the wage equation we retained only the variables U, ELT9, and F; in the union coverage equation we retained only lnW and BIGEMP; in the labor quality

TABLE 4

Estimated Regression Coefficient for the Union Coverage Variable from 20-industry sub-samples

A. Using the fully-specified model:

	Sample 1	Sample 2
OLS	.740**	.472**
	(.189)	(.193)
2 eqn. 2SLS	.703**	.674**
	(.222)	(.296)
2 eqn. 3SLS	.673	.622*
	(.221)	(.292)
3 eqn. 2SLS	.565*	.261
	(.309)	(.641)
3 eqn. 3SLS	.530**	.609
	(.216)	(.409)

B. Using the under-specified model:

	Sample 1	*Sample 2*	*Full sample*
OLS	.422*	.265*	.316**
	(.193)	(.124)	(.103)
2 eqn. 2SLS	.444*	.372*	.428**
	(.224)	(.162)	(.131)
2 eqn. 3SLS	.497*	.376*	.450**
	(.223)	(.163)	(.131)
3 eqn. 2SLS	.403	.387*	.385**
	(.265)	(.194)	(.138)
3 eqn. 3SLS	.471*	.423*	.451**
	(.260)	(.178)	(.134)

** Significant at the .01 level
* Significant at the .05 level

Standard errors are in parentheses. The tests are 1-tailed.

equation we retained only URBAN and OLD55. We estimated this underspecified model with our full 40 industry sample. Finally, we estimated the underspecified model for each of the 20-industry sub-samples. These results, for the union coverage coefficient only, are presented in Table 4. Our results seem to indicate that underspecification, by itself, is not sufficient to explain the instability of results obtained by Ashenfelter and Johnson, since our estimates of the underspecified equation for the full 40-industry sample remain large, fairly stable in value, and statistically significant as we move from OLS to 2-3SLS techniques. The estimates for our fully-specified model obtained from the 20-industry sub-samples, however, demonstrate considerably greater instability, with generally smaller t-values, falling to insignificance for one of the sub-samples as we move to 2-3SLS for the three-equation model. This would appear to indicate that it is the small degrees of freedom available in the Ashenfelter and Johnson study which led to their weak and unstable results. However, when we combine the problems of small sample size with underspecification of the model, we find the perplexing result that there is somewhat *less* instability in the value of the estimates, compared to the small-sample estimates obtained for the more fully-specified model, although significance levels are generally lower.

Our second supplementary estimation involved comparability with the previous Canadian studies by Kumar (1972) and Evans/MacDonald (1976) which used aggregate industry data. Both of these studies included a productivity variable in the wage equation, defined as value-added per manhour, and found it to be statistically significant. In an attempt to check whether our omission of this variable could explain the difference between our results and those of Kumar and Evans/MacDonald, we introduced it into our equation (1) as an additional variable. Results of OLS estimation indicated the value-added per man hour variable was nonsignificant (with a t value less than unity), and the coefficient of U was virtually unaffected by the addition of the productivity measure.

The third supplementary estimation entailed the estimation of various functional forms for the wage equation, a procedure suggested by Addison and Siebert (1979, pp. 271-2) as being reasonable. By and large these estimations confirmed that equation (1) was as good a functional form as any, but one variant suggested by Addison and Siebert, involving the addition of a U^2 term, produced interesting results:

$$
\begin{aligned}
(4)\ \ln W = {} & \underset{(.213)}{1.544^{**}} - \underset{(.396)}{.755U} + \underset{(.288)}{.900\,U^{2**}} - \underset{(.175)}{.578ELT9^{**}} - \underset{(.077)}{.513F^{**}} \\
& + .0112\,LAGPROF^{**} - \underset{(.144)}{.00344\,ONT} + \underset{(.142)}{.266\,QUE} \\
& + (.00393) \\
& + \underset{(.248)}{.448\,PRA} + \underset{(.192)}{.400\,BC^{*}}
\end{aligned}
$$

Standard errors are in parentheses; asterisks indicate significance levels as in table 1. The interesting aspect of equation (4) is that it yields an estimate of the union-nonunion differential of only 16 percent. This results because the equation indicates the union wage is *below* the nonunion wage until the industry becomes approximately 42 percent unionized. While such an effect is conceivable,[7] we have only two observations involving an industry with less than 42 percent union coverage. These results underpin our previous conclusion that our estimates of the union-nonunion differential are tenuous because the *nonunion* wage is not estimable with precision from our data set.

CONCLUSION

The initial motivation for this study was derived from a paper by Ashenfelter and Johnson (1972) in which they argue: 1) that the single wage equation model typically used to estimate the effect of union coverage on wages is inaccurate due to the presence of endogenous variables among the explanatory variables; and 2) that the estimated effect of union coverage on wages is small and nonsignificant when obtained from a more accurate simultaneous-equations model. Our results confirm point (1), but refute point (2). We do find strong evidence of joint dependence among the wage rate and two of its explanatory variables, union coverage and labor quality. However, allowance for this joint dependence in estimation does not reduce the estimated effect of union coverage on wages to nonsignificance. We continue to find a large and significant union wage effect. Our estimation results imply that the average wage for production workers in manufacturing industries is higher by 4.2 percent in response to a difference of 0.10 in union coverage (as measured by the proportion of all production workers covered by a collective bargaining agreement). Our results also imply an

6 This result is obtained as follows:

$$d(\ln W) = .412\, dU$$

or

$$\ln W^1 - \ln W^0 = .412\,(.10)$$
$$\ln(W^1/W^0) = .0412$$
$$W^1/W^0 = e^{.0412} = 1.0421$$
$$W^1 = 1.0421\, W^0$$

7 The "threat effect" may be small for low levels of unionization in an industry, allowing the "displacement effect" to more than offset (in the calculation of average union plus nonunion wages in the industry) the direct effect of unions on the wages of covered workers. See JOHNSON and MIESZKOWSKI (1970).

overall union-nonunion wage differential of 51 percent, but the accuracy of this figure depends critically on the accuracy with which we can estimate the average wage in an industry with no unionization. Because our sample includes no such industries and in fact contains only 3 (out of 40) industries with union coverage rates below 50 percent, we place little confidence in our estimate for the overall differential.

Additional estimation results were obtained in an effort to uncover the reasons for the disparate estimates of the union wage effect obtained in this and previous (U.S. and Canadian) studies. While these results were not clear-cut, it seems that differences in model specification, in industries selected for inclusion, and in sample size could all bear some responsibility. However, our results were fairly robust through all the experimental variations we tried. We suspect, though could not test with our data set, that a major cause of the disparity in estimates of the union wage effect obtained to date is the differing levels of aggregation on which the various estimates have been based.

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Data Appendix

The 40 industries in our sample, listed below, include some two, three, and four digit manufacturing industries. Our selection criterion was a pragmatic one; we included all those manufacturing industries for which the needed data were available.

List of Industries in the Sample

1970 SIC	*Name*
1011	Slaughtering and Meat Processing
102	Fish Products
104	Dairy Products
1071	Biscuit Manufacturers
1072	Bakeries
1081	Confectionery Manufacturers
1091	Soft Drink Manufacturers
1093	Breweries
15	Tobacco Products
162	Rubber Products
174	Shoe Factories
182	Wool, Yarn, and Cloth Mills
183	Man Made Fibre, Yarn, and Cloth Mills
23	Knitting Mills
243	Men's Clothing
245	Children's Clothing
251	Sawmills, Planing Mills and Shingle Mills
252	Veneer and Plywood Mills
271	Pulp and Paper Mills
273	Paper Box and Bag Manufacturers
289	Publishing and Printing
291	Iron and Steel Mills
294	Iron Foundaries
301	Boiler and Plate Works
302	Fabricated Structural Metal
304	Metal Stamping, Pressing and Coating
305	Wire and Wire Products Manufacturers
307	Heating Equipment Manufacturers
311	Agricultural Implements
315	Miscellaneous Machinery and Equipment Manufacturers
321	Aircraft and Aircraft Parts Manufacturers

323	Motor Vehicle Manufacturers
325	Motor Vehicle Parts and Accessory Manufacturers
327	Shipbuilding and Repair
331	Manufacturers of Small Electrical Appliances
332	Manufacturers of Major Appliances (Electric and Non-Electric)
334	Manufacturers of Household Radio and Television Receivers
336	Manufacturers of Electrical Industrial Equipment
36	Petroleum and Coal Products
375	Paint and Varnish Manufacturers

Notes on Sample Data

The data for profit and percent covered by collective agreements are reported by 1960 SIC, and were converted using the table in the 1970 Standard Classification Manual (12-501). However, since the SIC numbers were not specified and, for some industries, names did not exactly correspond, interpretive judgement was exercised.

Census data (for Education, Age and Urbanization) are published by 1970 SIC by 3 digit industry. Since some of the 1960 SIC industries become 4 digit industries in the 1970 SIC, data for those industries for the above variables were proxied with the appropriate 3 digit ratios. Further, census series combined industries 323 and 325; each of these is proxied by the ratio of the two combined.

Variable Definitions and Data Sources

U	Percent of Non-office Employees ("those non-supervisory workers directly engaged in the production of goods or services and the provision of maintenance and auxiliary services closely associated with production operations") covered by collective agreements; from Labour Canada, *Working Conditions in Canadian Industry 1974*. Surveyed are firms with 20 or more employees in the year ending 1 October 1974, with some exceptions covering smaller firms.
W	Wages Paid (gross earnings of employees before deductions for income tax and employee contributions to unemployment insurance, pensions, etc.) divided by manhours worked, for Production and Related Workers (those engaged in production and assembling activities, in storing, inspecting, handling, packing, warehousing, etc. and those engaged in maintenance, repair, janitorial and watchmen services, and line supervisors engaged in similar work to that of the employees they supervise); from Statistics Canada, *Manufacturing Industries of Canada, 1974*, 31-203.
ONT QUE	Number of Establishments located in Ontario, Quebec, The Prairie Provinces, and British Columbia, respectively, as a ratio to the total

Appendix Table: Data Used in Estimations (Rounded for presentation)

SIC70	*W*	*U*	*ELT9*	*F*	*LAGPROF*	*BIGEMP*	*URBAN*	*OLD55*	*ONT*	*QUE*	*PRA*	*BC*
1011	4.58	.86	.41	.17	2.3	.59	.73	.13	.37	.28	.22	.09
104	4.09	.69	.34	.09	3.7	.32	.51	.14	.36	.26	.21	.05
102	3.11	.73	.60	.39	8.7	.46	.11	.15	.04	.12	.05	.13
1071	3.47	.87	.39	.59	2.6	.71	.70	.13	.41	.32	.15	.07
1072	3.69	.58	.39	.30	2.6	.33	.70	.13	.31	.27	.21	.15
1081	3.27	.79	.35	.57	5.3	.76	.75	.13	.43	.33	.11	.09
1091	4.04	.59	.28	.07	5.9	.19	.74	.12	.30	.37	.14	.06
1093	5.89	.98	.28	.00	10.0	.71	.74	.12	.25	.07	.36	.14
15	5.04	.95	.38	.44	8.8	.96	.73	.14	.38	.63	.00	.00
162	4.37	.90	.32	.18	5.6	.85	.63	.11	.60	.23	.05	.11
174	2.79	.57	.53	.62	3.7	.46	.69	.13	.43	.51	.04	.02
182	3.26	.82	.48	.38	7.4	.72	.36	.16	.38	.42	.04	.02
183	3.42	.75	.39	.27	6.3	.75	.21	.13	.41	.54	.01	.01
23	2.70	.38	.50	.65	5.1	.43	.62	.13	.28	.66	.02	.03
243	2.79	.60	.56	.78	4.4	.44	.72	.14	.22	.62	.12	.03
245	2.56	.17	.55	.81	3.2	.10	.85	.13	.16	.82	.02	.00
251	4.75	.73	.48	.01	13.9	.40	.20	.14	.16	.32	.08	.25
252	4.75	.84	.37	.15	14.0	.66	.49	.11	.27	.33	.04	.33
271	5.62	.97	.33	.02	7.0	.95	.33	.14	.25	.38	.07	.16
273	4.29	.76	.32	.22	7.0	.48	.83	.10	.51	.29	.08	.07
289	5.41	.52	.11	.19	14.1	.51	.76	.12	.37	.08	.32	.13
291	5.59	.75	.32	.01	10.9	.95	.65	.12	.36	.36	.13	.11
294	4.44	.88	.47	.02	5.4	.57	.53	.17	.51	.23	.10	.11
301	5.62	.91	.29	.00	5.3	.56	.73	.14	.44	.23	.14	.06
302	5.43	.89	.33	.00	5.0	.64	.83	.12	.44	.21	.15	.14
304	4.38	.68	.34	.13	9.1	.34	.80	.11	.55	.23	.11	.09
305	4.54	.82	.32	.12	9.9	.46	.78	.12	.59	.25	.05	.10
307	4.01	.68	.29	.06	5.2	.31	.83	.12	.65	.18	.10	.06
311	4.95	.89	.34	.01	6.0	.78	.59	.18	.38	.12	.45	.03

Appendix Table: Data Used in Estimations (Rounded for presentation)

SIC 70	*W*	*U*	*ELT9*	*F*	*LAGPROF*	*BIGEMP*	*URBAN*	*OLD55*	*ONT*	*QUE*	*PRA*	*BC*
315	4.69	.81	.24	.05	5.6	.49	.68	.11	.58	.18	.10	.12
321	4.76	.81	.22	.06	3.3	.84	.91	.17	.36	.27	.18	.15
323	5.90	.99	.29	.01	6.0	.98	.66	.09	.45	.14	.05	.23
325	4.94	.91	.29	.19	6.0	.82	.66	.09	.76	.10	.06	.07
327	5.03	.93	.43	.00	.1	.81	.63	.23	.19	.10	.02	.38
331	3.60	.88	.27	.51	9.7	.79	.72	.12	.78	.14	.04	.04
332	4.15	.92	.33	.11	4.7	.91	.75	.13	.74	.19	.00	.00
334	3.48	.84	.24	.56	3.7	.93	.81	.08	.71	.29	.00	.00
336	4.12	.90	.21	.25	5.3	.70	.68	.12	.58	.26	.06	.09
36	6.18	.84	.16	.00	11.9	.61	.74	.12	.30	.24	.27	.10
375	4.29	.42	.20	.07	5.4	.31	.90	.14	.49	.26	.07	.14
Mean	4.35	.77	.35	.23	6.5	.62	.66	.13	.42	.30	.11	.10

PRA BC	number of establishments in Canada; from Statistics Canada, 31-203, 1974. Establishment is defined as the smallest unit which is a separate operating unit capable of reporting certain principal statistics; typically a factory, mill or plant.
ELT9	The number of employees with completed education less than grade 9, divided by the total number of employees; from Statistics Canada 1971 Census publication, *Industries by Sex Showing Age, Marital Status, Level of Schooling and Class of Worker for CANADA*, No. 94-749.
BIGEMP	Ratio of the number of employees in establishments employing 200 or more employees divided by the total number of employees; from Statistics Canada, *Type of Organization and Size of Establishments, 1974,* 31-210.
F	Proportion of production and related employees that are female; from Statistics Canada 31-203.
LAGPROF	Percent profit before tax on total income excluding investment income; from Statistics Canada, *Corporation Financial Statistics*, 1974, 61-207.
URBAN	Labour force residing in Census Metropolitan Areas divided by total Labour Force. Labour force residing in CMA's is from Statistics Canada 1971 Census publication *Industries by Sex for Census Metropolitan Areas, Place of Residence and Place of Work* 94-742; total labour force is from 1971 Census publication *Industries by Sex for Canada, Regions and Provinces,* 94-740.
OLD55	Employed labour force age 55 years and over divided by the total employed labour force; from Statistics Canada, 1971 Census publication, *Employed Labour Force by Industry, Age and Sex for Canada and Provinces,* 94-747.

Le syndicalisme et les salaires: un réexamen de la théorie

La raison première de cet article provient d'une étude d'Ashenfelter et de Johnson (1972) dans laquelle ils énonçaient ce qui suit: 1° le modèle simple d'équation salariale, typiquement utilisée pour mesurer l'effet des syndicats sur les salaires, est inexact à cause de la présence de variables endogènes parmi les variables explicatives; 2° l'effet présumé des syndicats sur les salaires est de peu d'importance et sans signification quand il est obtenu à partir d'un modèle d'équation simultanée plus précis. Les résultats de notre étude confirment la première proposition, mais ils démontrent la fausseté de la seconde.

Nous avons trouvé une preuve solide de dépendance réciproque entre le taux des salaires et deux de ses variables explicatives, soit la présence de syndicats et la qualité du travail. Cependant, le fait de tenir compte de cette double dépendance ne réduit pas l'effet présumé de l'existence des syndicats à néant. Au contraire, nous persistons à déceler que les salaires conventionnels ont un effet considérable et

significatif. Les résultats de notre évaluation indiquent que le salaire moyen des travailleurs de la production dans les industries manufacturières est supérieur à 4.2 pour cent en regard d'une différence de 0.10 dans le degré de pénétration syndicale (celle-ci étant la proportion de tous les travailleurs de la production liés par une convention collective de travail). Les résultats que nous avons obtenus indiquent aussi une différence de 51 pour cent entre les salaires conventionnels et les autres, mais l'exactitude de cette estimation dépend, et cela d'une façon décisive, de l'exactitude avec laquelle nous pouvons établir le salaire moyen dans une industrie où il n'y a pas de syndicat. Parce que notre échantillon n'inclut pas de telles industries et ne contient en réalité que trois industries (sur 40) où le taux de syndicalisation est inférieur à 50 pour cent, nous avons peu de confiance dans notre estimation pour la différence d'ensemble.

Les résultats d'une évaluation supplémentaire ont été obtenus dans un effort pour découvrir les raisons d'estimations disparates de l'effet du salaire conventionnel dans la présente étude et dans des études antérieures aux États-Unis et au Canada. Quoique ces résultats ne soient pas bien précis, il semble que les différences dans la spécification du modèle, les industries sélectionnées et l'ampleur de l'échantillon peuvent tous en être responsables. Cependant, les résultats auxquels nous sommes arrivés étaient solides dans toutes les expérimentations variées que nous avons tentées. Nous soupçonnons, bien que nous n'ayons pu le vérifier à partir des données dont nous disposions, que la cause principale de la disparité dans les estimations de l'effet du salaire conventionnel obtenues jusqu'ici réside dans les différents niveaux d'agrégation sur lesquels les diverses évaluations ont été fondées.